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STANDARD ERROR OF AN EQUATING BY
ITEM RESPONSE THEORY

Frederic M. Lord

This research was sponsored in part by the
Personnel and Training Research Programs
Psychological Sciences Division
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Contract Authority Identification Number
NR No. 150-453

Frederic M. Lord, Principal Investigator



Educational Testing Service
Princeton, New Jersey

November 1981

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A formula is derived for the asymptotic standard error of a true-score equating by item response theory. The equating method is applicable when the two tests to be equated are administered to different groups along with an 'anchor test.' Numerical standard errors are shown for an actual equating (1) comparing the standard errors of IRT, linear, and equipercentile methods; (2) illustrating the effect of the length of the anchor test on the standard error of the equating. ←		

Standard Error of an Equating by Item Response Theory

Abstract

A formula is derived for the asymptotic standard error of a true-score equating by item response theory. The equating method is applicable when the two tests to be equated are administered to different groups along with an 'anchor test.' Numerical standard errors are shown for an actual equating 1) comparing the standard errors of IRT, linear, and equipercentile methods; 2) illustrating the effect of the length of the anchor test on the standard error of the equating.

Standard Error of an Equating by Item Response Theory*

In item response theory (IRT), an examinee's expected number-right score ξ on test X is equal to the test characteristic function evaluated at the examinee's ability level θ :

$$\xi = \sum_{g=1}^{n_x} P_g(\theta) \quad (1')$$

where $P_i(\theta)$ is the item response function, the probability of a correct answer to item i at ability level θ . If we have a second test, Y , measuring the same ability as X , the expected number-right score η on this test may be written as

$$\eta = \sum_{h=1}^{n_y} P_h(\theta) \quad (4')$$

Equations (1') and (4') are parametric equations for the functional relationship between ξ and η . Note that this relationship is an exact mathematical one, not a statistical association. Given any θ , (1') and (4') determine a pair of values, ξ and η , that represent the same ability level as θ . Pairs of values (ξ, η) determined in this way are equated. In practice, it is often assumed that the functional relationship of η to ξ given by (1') and (4') can also be applied to actual number-right scores on the two tests, producing an equating of these scores.

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Here, we simply deal with the sampling errors in estimating the equating relationship of η to ξ . In (1') and (4'), estimated item parameters must be used. These are the source of the sampling errors in IRT equating. Note that the ability estimates for individual examinees are not used in (1') and (4') and thus will not appear in our formulas. Until now, the sampling errors of IRT equatings have never been estimated.

Data

In IRT equating, we frequently have a set of common items that are administered to all examinees. These are needed in order to get Test Y item parameters on the same scale as Test X item parameters. If the common items are external to tests X and Y, as assumed here, the common items are called the anchor test, or, in the present report, Test W. The sampling variance formulas to be obtained here can be modified in obvious ways for the case where some or all of the common items are internal to the tests that are being equated.

Designate the examinees who took both Tests X and W as Group 1; designate the examinees who took Tests Y and W as Group 2. Typically, every examinee falls in one of these two groups.

In practice when there is a series of test forms A, B, ..., X, Y, Z, ... (say), the 'Group 1' data on Test X are processed as soon as they become available in order to equate Test X to the preceding form. When the Group 2 data become available at some later date, it is often considered uneconomical to rerun the Group 1 data, so Group 2 is

run by itself. This case, where item parameters for Groups 1 and 2 are estimated separately, is the case to be considered here. (The simplifying assumption that is used below to approximate the sampling variances of the estimated item parameters is not available in the alternative case where Groups 1 and 2 are pooled and all parameters estimated simultaneously.)

New Equating Formulas

When parameters are estimated separately for groups 1 and 2, the item parameters and θ in (4') have a different origin and scale from the item parameters and θ in (1'). It is thus no longer possible simply to eliminate θ from (1') and (4') to obtain the relation of η to ξ . The customary procedure in this situation is to use the anchor test to transform the Group 2 item parameters on to the scale of the Group 1 item parameters. This procedure adds to the sampling variance of the transformed item parameters and greatly complicates any determination of the sampling variance of the subsequent equating. The procedures and formulas given below avoid this problem since they avoid any transformation of item parameters.

Equations (1') and (4') remain unchanged except that additional subscripts (explained below) are used. In particular, the symbols θ_1 and θ_2 must be distinguished because groups 1 and 2 use different ability scales:

$$\xi = \sum_g P_{gl}(\theta_1) \quad , \quad (1)$$

$$\eta = \sum_g P_{gp} g_4(\theta_2) \quad . \quad (4)$$

The item response functions here are written P_{gp} where $p = 1, 2, 3, 4$ refers to (test X, group 1), (test W, group 1), (test W, group 2), and (test Y, group 2) respectively, and $g = 1, 2, \dots, n_p$ where n_p is the number of items in the appropriate test.

Let us write down similar equations for the expected number-right score ω on anchor test W :

$$\omega = \sum_g P_{gp} g_2(\theta_1) \quad , \quad (2)$$

$$\omega = \sum_g P_{gp} g_3(\theta_2) \quad . \quad (3)$$

The equation numbering keeps the tests in convenient order. The desired equation relation between η and ξ can be obtained by eliminating θ_1 , θ_2 , and ω from these four equations.

Computer programs are available for equating η to ξ by eliminating θ from (1') and (4'). These same programs can be used to equate ω to ξ in one step, using (1) and (2), then to equate η to ω in a second step using (3) and (4). This produces an equating of η to ξ for the presently relevant situation where Group 1 and Group 2 parameters are not on the same scale.

An estimated equating is obtained from (1) - (4) after replacing the true item parameters by their maximum likelihood estimates. Using carets to denote this change, we have

$$\xi = \sum_g \hat{P}_{g1}(\theta_1) \quad , \quad (1'')$$

$$\hat{\omega} = \sum_g \hat{P}_{g2}(\theta_1) \quad , \quad (2'')$$

$$\hat{\omega} = \sum_g \hat{P}_{g3}(\theta_2) \quad , \quad (3'')$$

$$\hat{\eta} = \sum_g \hat{P}_{g4}(\theta_2) \quad . \quad (4'')$$

These equations show that $\hat{\eta}$ is a function of all the estimated item parameters together with the specified value of ξ .

Derivatives

For item g , instead of using a_g , b_g , and c_g to denote the three parameters commonly used in IRT, let us use t_{1gp} , t_{2gp} , and t_{3gp} , respectively. We will need certain derivatives for $r = 1, 2, 3$, obtained from (1'')-(4''):

$$\frac{\partial \hat{\eta}}{\partial t_{rg4}} = P_{g4}^{(r)}(\theta_2) \quad , \quad (5)$$

$$\frac{\partial \hat{\omega}}{\partial t_{rg3}} = P_{g3}^{(r)}(\theta_2) \quad ,$$

$$\frac{\partial \hat{\omega}}{\partial t_{rg2}} = P_{g2}^{(r)}(\theta_1) \quad ,$$

where $P_{gp}^{(r)}$ denotes the derivative of P_{gp} with respect to t_{rgp} .

Similarly,

$$\frac{\partial n}{\partial \theta_2} = \frac{\sum p' g_4(\theta_2)}{g} ,$$

$$\frac{\partial w}{\partial \theta_1} = \frac{\sum p' g_2(\theta_1)}{g}$$

where p' denotes a derivative with respect to θ . Using the formula for the derivative of an implicit function, we also find from (1")-(4") for $r = 1, 2, 3$

$$\frac{\partial \theta_2}{\partial t_{rg3}} = - \frac{p_{g3}^{(r)}(\theta_2)}{\sum p' g_3(\theta_2)} ,$$

$$\frac{\partial \theta_1}{\partial t_{rg1}} = - \frac{p_{g1}^{(r)}(\theta_1)}{\sum p' g_1(\theta_1)} ,$$

$$\frac{\partial \theta_2}{\partial w} = \frac{1}{\sum p' g_3(\theta_2)} .$$

Using the chain rule for derivatives, we find from the above formulas:

$$\frac{\partial n}{\partial t_{rg3}} = \frac{\partial n}{\partial \theta_2} \frac{\partial \theta_2}{\partial t_{rg3}} = - p_{g3}^{(r)}(\theta_2) \frac{\sum p' g_4(\theta_2)}{\sum p' g_3(\theta_2)} , \quad (6)$$

$$\frac{\partial \eta}{\partial t_{rg2}} = \frac{\partial \eta}{\partial \theta_2} \frac{\partial \theta_2}{\partial w} \frac{\partial w}{\partial t_{rg2}} = p_{g2}^{(r)}(\theta_1) \frac{\sum_p P'_{g4}(\theta_2)}{\sum_g P'_{g3}(\theta_2)} , \quad (7)$$

$$\frac{\partial \eta}{\partial t_{rg1}} = \frac{\partial \eta}{\partial \theta_2} \frac{\partial \theta_2}{\partial w} \frac{\partial w}{\partial \theta_1} \frac{\partial \theta_1}{\partial t_{rg1}} = -p_{g1}^{(r)}(\theta_1) \frac{\sum_p P'_{g2}(\theta_1)}{\sum_g P'_{g1}(\theta_1)} \frac{\sum_p P'_{g4}(\theta_2)}{\sum_g P'_{g3}(\theta_2)} . \quad (8)$$

Given ξ , we are now in a position to express $\hat{\eta}$ as a series in powers of $\hat{t}_{rgp} - t_{rgp}$ ($r = 1, 2, 3$; $g = 1, 2, \dots, n_p$; $p = 1, 2, 3, 4$). We will write η'_{rgp} instead of $\partial \eta / \partial t_{rgp}$ and η''_{rgpshq} instead of $\partial^2 \eta / \partial t_{rgp} \partial t_{shq}$:

$$\begin{aligned} \hat{\eta} = \eta + \sum_p \sum_g \sum_r (\hat{t}_{rgp} - t_{rgp}) \eta'_{rgp} \\ + \frac{1}{2} \sum_p \sum_q \sum_g \sum_h \sum_r \sum_s (\hat{t}_{rgp} - t_{rgp}) (\hat{t}_{shq} - t_{shq}) \eta''_{rgpshq} + \dots . \quad (9) \end{aligned}$$

Sampling Variance

Transposing, squaring, and taking expectations, we find from

(9) for fixed ξ ,

$$\text{Var} \hat{\eta} = \delta (\hat{\eta} - \eta)^2 = \sum_p \sum_q \sum_g \sum_h \sum_r \sum_s \eta'_{rgp} \eta'_{shq} \text{Cov}(\hat{t}_{rgp}, \hat{t}_{shq}) + \dots .$$

When item parameters and abilities are both estimated simultaneously by maximum likelihood, it is not practical to use the usual sampling covariance formulas for all estimators simultaneously. As a rough approximation, it is customary (Lord, 1980, Section 12.3) to use instead the (simpler) formulas for the case where the ability parameters are known. We will use this rough approximation here to find $\text{Cov}(\hat{t}_{rgp}, \hat{t}_{shq})$. Because of this approximation, our sampling variance of equating will be an underestimate.

In this case, all covariances involving two different items are exactly zero, as are all covariances involving a single item administered to two different groups of examinees. All nonzero variances and covariances are inversely proportional to N , the number of examinees.

We now have

$$\begin{aligned} \text{Var} \hat{n} = & \sum_{pg} \left[\sum_{r=1}^3 \sum_{s=1}^3 \{ \eta'_{rgp} \eta'_{sgp} \text{Cov}(\hat{t}_{rgp}, \hat{t}_{sgp}) \} \right. \\ & \left. + \sum_{r=1}^3 \sum_{s=1}^3 \{ \} + \sum_{r=1}^3 \sum_{s=1}^3 \sum_{t=1}^3 \{ \} + \dots \right] . \end{aligned}$$

Some higher order terms are indicated here in order to make clear that the number of terms under summation signs does not increase too rapidly. The triple summation represents 3 times as many terms as the double summation, but each term in the triple summation is divided by $N^{3/2}$ whereas each term in the double summation is only divided by N . When N is several thousand, it is reasonable to expect that the higher order terms can be neglected, as is customary with asymptotic variances.

Our final asymptotic formula, then is

$$\text{Var } \hat{n} = \sum_{p=1}^4 \sum_{g=1}^{n_p} \sum_{r=1}^3 \sum_{s=1}^3 n'_{rgp} n'_{sgp} \text{Cov}(\hat{t}_{rgp}, \hat{t}_{sgp}) . \quad (10)$$

The n' values required here are computed from (5) - (8). The covariances are obtained by the usual formulas for covariances of maximum likelihood estimators of item parameters when ability parameters are fixed (Lord, 1980, p. 191).

Practical Application

Without data, it is difficult to make inferences about the magnitude of the sampling errors in IRT equating. Will they be larger or smaller than the sampling errors in conventional linear equating? In conventional equipercentile equating? Do sampling errors become large or small at extreme score levels?

Equation (10) has been applied to an equating of the Verbal score on the 90-item Form VSA4 of the Scholastic Aptitude Test (12/73 administration) to the 85-item Form XSA2 Verbal score (4/75 administration). All examinees took an SAT and also a 40-item anchor test. Petersen, Cook, and Stocking (1980) made separate LOGIST runs on the 130 items in the 1973 administration for a sample of 2665 examinees, and on the 125 items in the 1975 administration for a sample of 2686 examinees. They have allowed the use here of their item parameter estimates.

SAT scaled scores are a linear transformation of formula scores (rights minus one-quarter wrongs). Our results here are for the hypothetical case where all examinees answer all items. In this special case formula scores are a linear transformation of number-right scores, so scaled scores are likewise. Since a known linear transformation $A\xi + B$ of number-right scores ξ simply multiplies the standard error of ξ by the constant A , it is not difficult to obtain scaled-score standard errors from (10). A computer program to do this was written and run by Marilyn Wingersky.

For each of certain specified formula scores on XSA2, Table 1 shows 1) the equivalent scaled score found by the conventional linear procedure usually used for the SAT (Design IV A, Angoff, 1971), 2) the standard error of these equated (scaled) scores as found by the computer program AUTEST (Lord, 1975) assuming the validity of the linear model; also 3) the equivalent scaled score found by the IRT method of this report, and 4) the corresponding scaled-score standard error calculated from (10). The standard errors in Table 1 are best understood in comparison with the standard deviation of scaled scores, which is 106 for XSA2; and in comparison with the classical test theory standard error of measurement (due to imperfect test reliability), which is 31. Clearly the standard error of equating is small compared to the standard error of measurement.

Judging by the IRT standard errors, the equating is definitely nonlinear, at least outside the score range from 350 to 650. The IRT standard errors show a continued sharp increase as the minimum

Table 1

A Comparison of Linear and IRT Equatings and of Their Standard Errors

Selected formula scores*, XSA2	Linear Model		IRT Model	
	Equivalent scaled score	Standard error	Equivalent scaled score	Standard error
84	780	4.6	813.8	2.3
79.74	750	4.2	778.0	4.5
72.70	700	3.6	717.6	4.4
65.65	650	3.1	658.8	3.6
58.61	600	2.5	602.4	2.8
51.57	550	2.1	548.0	2.2
44.52	500	1.7	495.4	2.0
37.48	450	1.5	445.7	2.1
30.43	400	1.6	399.3	2.3
23.39	350	1.8	355.6	2.8
16.35	300	2.3	313.3	3.6
9.30	250	2.8	270.2	4.7
2.26	200	3.3	223.0	7.0
-5	150	3.9	163.5	15.6

*Although formula score is actually a discrete variable, it is for convenience treated here as continuous.

possible true formula score of -5.5 is approached. At the other end of the score scale, the IRT standard error increases up to a scaled score of 760 and decreases thereafter. The reason for the decrease at the upper end is that for a perfect score, the standard error of this kind of IRT equating is zero. Except at the upper end, the IRT standard error is larger than the linear.

The results of Table 1 are displayed in Figures 1-2. The straight line in Figure 1 shows the linear equating of true formula score on XSA2 to true scaled score on VSA4. The dashed lines are drawn two standard errors above and below the straight line.

Figure 2 similarly displays the curvilinear IRT equating of XSA2 to VSA4 and its standard error. The straight-line extension of the lower end of the equating (middle) line in Figure 2 was obtained by the method described in Lord (1980, pp. 210-211). It is shown in the figure for completeness, but no standard error is shown since there is no good theoretical basis for such an extension.

Table 2 compares present IRT equating with a conventional equipercentile equating of XSA2 to VSA4 via the anchor test. In conventional equating, an XSA2 score and a VSA4 score each equipercentile-ly equivalent to a given anchor test score are taken to be equivalent to each other. The standard error of the resulting equipercentile equating of XSA2 to VSA4 is given by $\sqrt{(SE_{XSA2}^2 + SE_{VSA4}^2)}$ where the SE under the radical sign are standard errors of separate equipercentile equatings of each test to the anchor test. Formulas for SE_{XSA2} and SE_{VSA4} are given in Lord (1981).

Figure 1. Linear equating of true formula score on XSA2 to true scaled score on VSA4. Dashed lines are two scaled-score standard errors above and below equating line.

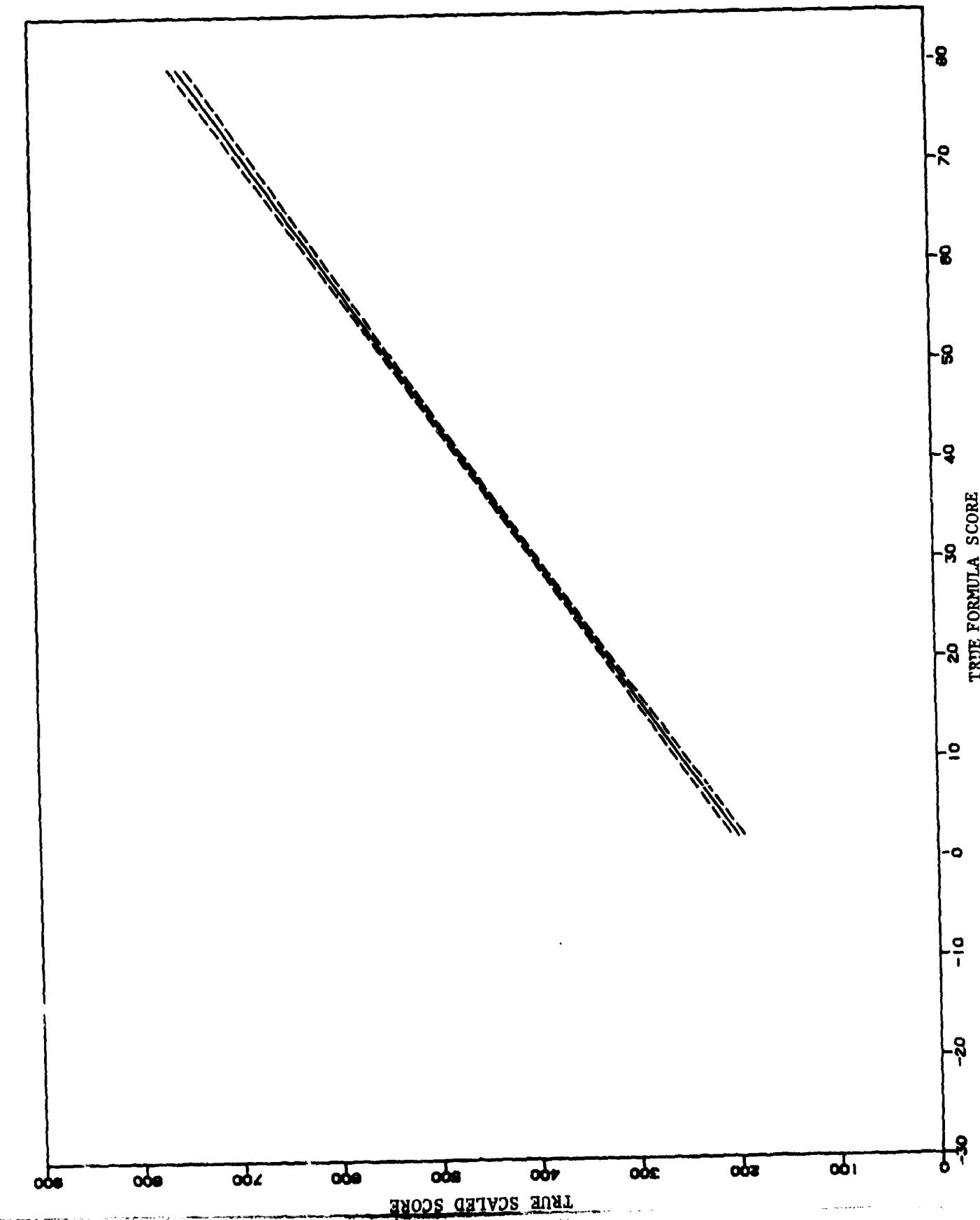


Figure 2. IRT equating of XSA2 formula score to VSA4 scaled score, with two-standard-error bounds.

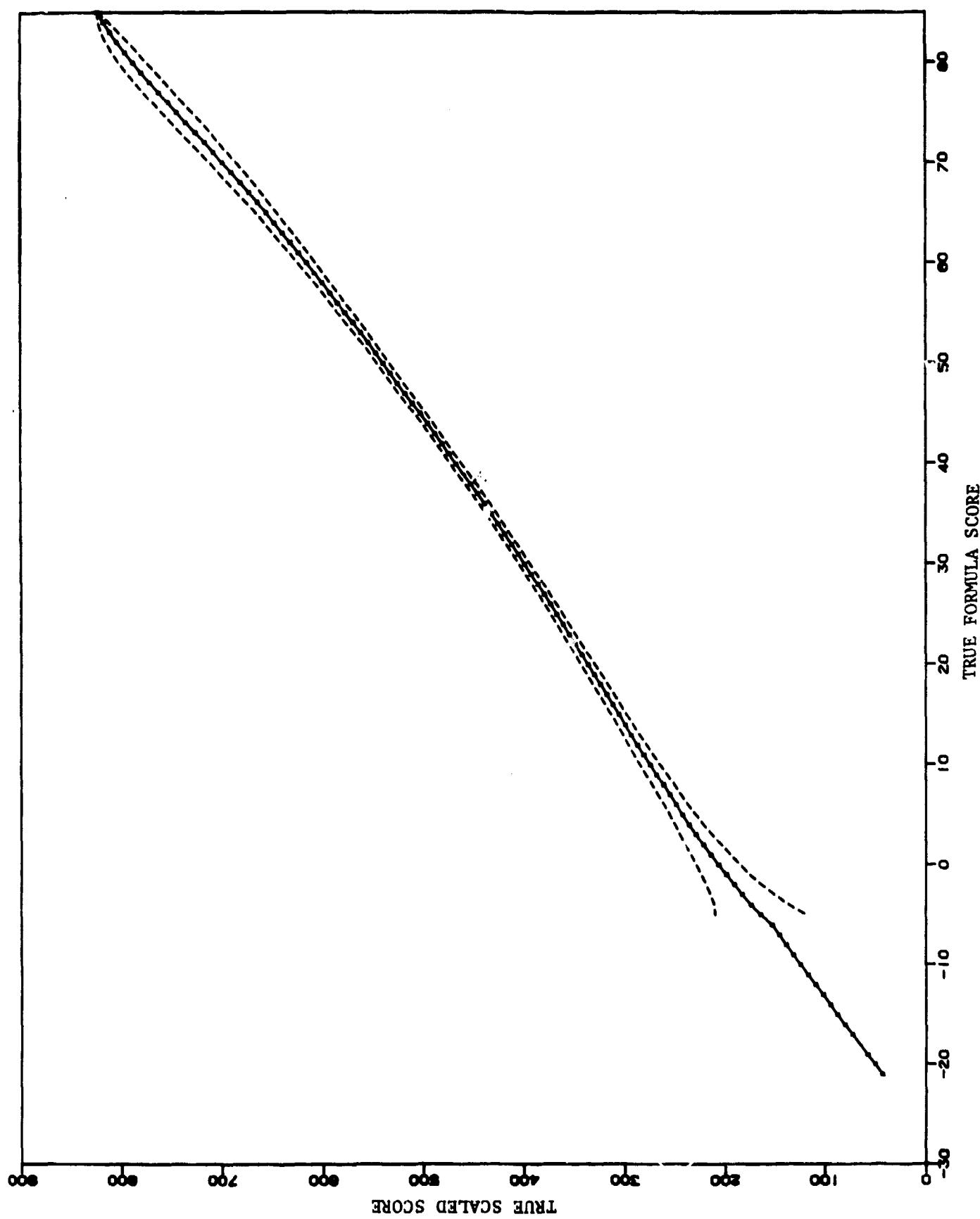


Table 2

A Comparison of Equipercentile and IRT Equating
and of Their Standard Scores

XSA2 formula score	Equipercentile Method		IRT Model	
	Equivalent scaled score	Standard error	Equivalent scaled score	Standard error
78.1	774	13.47	764	4.68
70.6	722	15.85	700	4.18
64.75	652	10.32	651	3.44
58.9	602	4.97	605	2.78
52.9	558	4.12	558	2.32
47.25	514	3.47	515	2.09
40.1	466	3.44	464	2.05
32.4	417	2.93	412	2.24
25.75	364	3.37	370	2.63
16.1	314	4.07	312	3.62
7.6	242	5.70	259	5.08
-3.75	195	7.85	175	12.49

Since SE_{XSA2} and SE_{VSA4} are estimated from unsmoothed data, the equipercentile standard errors in Table 2 fluctuate somewhat. Nevertheless, it is apparent that the equipercentile method has a much larger standard error above a scaled score of 450. For these data, the IRT method shows a larger standard error than the equipercentile method only when the formula score is negative.

The standard error of equipercentile equating could be reduced by smoothing the frequency distribution of raw scores before equating. Smoothing is undoubtedly desirable as a practical expedient; however the choice of a smoothing formula is somewhat arbitrary and the smoothing is likely to prevent convergence of the estimated equating to its true value in large samples. Formulas for the standard errors of smoothed equipercentile equating are not presently available.

In order to determine the effect of using a shorter anchor test, every other item in the anchor test was discarded and the data reanalyzed on the basis of the remaining 20-item anchor test. The effect on the standard errors of IRT equating is shown in Table 3. The two equatings agree fairly well. At the point where the equating standard errors are a minimum, halving the length of the anchor test increases the standard error by a factor of about $\sqrt{2}$. At the other score points, the effect is less. Given standard errors like those in Table 2, it will now be possible to make a reasonable judgment as to the length necessary for an anchor test.

Table 3

IRT Equatings and Their Scaled-Score Standard Errors,
a Comparison of Results Using 20- and 40-Item Anchor Tests

XSA2 formula score	Length of Anchor Test			
	20 Items		40 Items	
	Scaled score	Standard error	Scaled score	Standard error
80	787	5.9	780	4.5
70	698	5.3	695	4.1
60	615	3.9	613	2.9
50	540	3.0	536	2.2
40	467	2.7	463	2.0
30	399	3.0	397	2.4
20	336	3.9	335	3.2
10	274	5.4	275	4.6
0	206	9.9	206	8.4

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